Agri-Commodity Price Dynamics: The Relationship Between Oil and Agricultural Market

Franco Rosa, Michela Vasciaveo

Department of Food Science – University of Udine – Italy

franco.rosa@uniud.it, michela.vasciaveo@uniud.it

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Abstract

The aim of this paper is to analyze the interactions among the prices of some agricultural commodities in Italy and United States by using the time series analysis method. After a general overview of the world and European agri-markets, the agricultural commodity and oil prices are investigated in order to analyze the cross-market interactions and test the hypothesis that the increased volatility in agricultural prices is caused by the exogenous crude oil prices. For the analysis the data about the commodity spot price series of wheat, corn, soybeans in US and Italy and crude oil price are collected. The results suggest: i) the presence of causal nexus with an exogenous influence of the oil price on the agricultural commodities for the US markets; ii) the evidence of cointegration between US and Italian commodities supporting the unique price condition; iii) no clear evidence of causality between oil and Italian agri-commodities, suggesting that the oil volatility is transmitted directly to the US market and indirectly to the Italian one.
Key words: time series analysis, agricultural commodity prices, volatility, causality, market integration.

1. Introduction

Price regime

Since late 2007, the major grains and oilseeds markets have experienced a period of tight supplies and strong demand causing higher price volatility rising sharply in 2007, followed by a decline in early 2008[1]. The price spike spanning the 07-08 period seems to be affected also by a macroeconomic scenario influencing the world commodity markets as the rapid growth in food demand by the BRIC countries due to rise in GDP, the international financial crisis, and growing influence of the oil prices on the commodity markets (Piot-Lepetit and M'Barek, 2011). The globalization and growing integration of financial and energy markets with agricultural commodity markets, has generated complex interactions among these markets and growing difficulties in understanding the price movements to be used in planting decisions. With the agrofuel commodities there is an additional so-called knock-on effect: the expanding U.S. corn production for ethanol reduces oilseed acreage, such that oilseed prices increase as a result of expected tightening supplies and this price strength is enhanced by the rising demand for both, meals being the cereal feed substitute and vegetable oils for bio-diesel production (OECD-FAO Agricultural Outlook 2007-2016).

[1] Wheat prices (SRW No. 2, f.o.b., U.S. Gulf ports) rose 58% between November 2007 and March 2008 - from $292,29 per ton to $461,86 before falling back in April and May. Maize prices (Yellow, No. 2, f.o.b., U.S. Gulf ports) rose 78% between November 2007 and June 2008, from $170,37 per ton to $303,24. Soybeans prices (Yellow No. 1, f.o.b.,U.S., Gulf ports) increased 65% passing from $380,57 per ton in November 2007 to $629,41 per ton in July 2008 (source USDA); in the same period the peak level of oil prices reached the $140 per barrel.

The movements of agricultural commodity prices are expected to be influenced by the following general factors:

- Growth in world population, incomes and food consumption with rising demand of staples from China and India due to new consumption patterns and demand growth for meat, catching the demand for grain and oilseeds used to produce meat (Headey and Fan, 2008);
- Change in stocks, stock replenishment rates and consequent effects on international trader expectations and decisions (Trostle, 2008). Rising commodity prices are reflective also of higher inflation costs and increased worldwide demand for food. Investors can buy grain stocks by buying commodity funds or commodity stocks producing grain stocks, or stocks that are needed to produce grain stocks.
- Growing speculation in financial markets, responsible of the increasing agricultural commodity prices in 2007–08 and leading to unreasonable or unwanted price fluctuations (Robles *et al.*, 2009). The excess price surges caused by speculation could cause severe impact on the global grain markets and, consequently, prevent the market's performance in responding to fundamental changes in supply, demand, and costs of production.
- Exchange rates fluctuations affecting the relative prices: price fluctuations for soft commodities can destabilize real exchange rates and cause difficulties for the governments in preserving a stable economic environment. Fluctuations in agricultural commodity prices create substantial risk for producers, processors and traders in managing revenues and the cost of future production and consumption. Prices are influenced by external factors, including the changes in domestic supply and demand, international prices, exchange rates and climatic, economic and political conditions. The higher and more unpredictable is the price volatility of agricultural commodities, the greater is the possibility of incurring losses (or realizing gains) on future transactions (sales or purchases) of those commodities. Moreover, the uncertainty reduces the opportunities to access credit markets and drives farmers to use lower risk production techniques (John, 2007).
- Sudden climate changes with impact on agricultural yields in different parts of the globe: most of the cereal production decline is due to reduced plantings and adverse weather in some major producing and exporting countries. In addiction, biofuel policies and their acceptance from traders (Sarris, 2009), policies encouraging the biofuel production from agricultural commodities, rising oil prices, limited supplies of fossil fuel, and increased concerns for global warming, have influenced the growing demand for renewable energy sources, and have prompted the surge of biofuel demands. This has justified the raise in corn price and the acreage invested in biofuel crops (Moschini and Hennessy, 2001) and reduction in planting soybeans,

that is in competition with corn with a significant surge in price up to 75% from April 2007 to April 2008 (Headey and Fan, 2008).

European overview

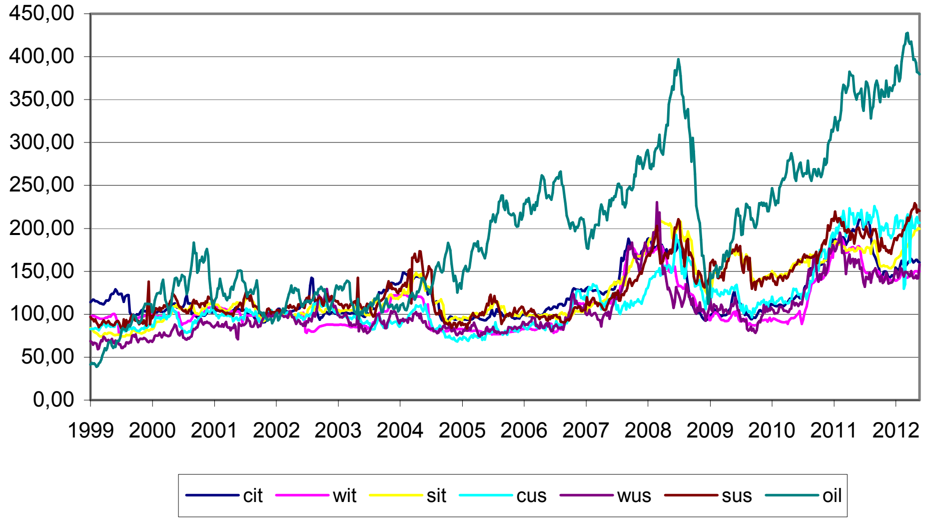
The European Union (EU) expanded from 15 to 25 countries in 2004; this was the largest single enlargement in terms of people, land and number of countries, and further broadened to 27 in 2007 with Bulgaria and Romania. The result is the EU population grew by 28 percent, the arable land increased by nearly 40 % and the number of countries almost doubled. Actually the EU generally accounts for about 15-20 percent of the world's agricultural exports and imports and the EU-27 is one of the most important trading partner and competitor of the United States in the world agricultural markets. European agricultural policy has long had a major impact on world agricultural markets, and the EU is one of the key participants in World Trade Organization (WTO) negotiations on agricultural trade.

In Italy agri-commodity markets are influenced by the Common Agricultural Policy (CAP) that evolved considerably during the past 20 years, moving from market support (first pillar) to the rural development policy beginning in 2003 (second pillar). The amount of intervention in cereal markets began to decline in 1992 with the Mac-Sharry Reform (-30%), and continued with the Agenda 2000 reform (-15%), showing non significant changes over the next years observed. The EU protection mechanism for cereals, even after the URAA (Uruguay Round Agreement Act) converted all border measures into import duties, for a long period resulting in a wide gap between entry (border) and intervention (domestic) prices (Listorti, 2009).

In 2003 the horizontal regulation (1782/2003) caused two effects: the first one maintained the previous set aside regime starting in 1992 with cereal and oleaginous reform, the second one supported the agri-energy crops with different provisions used to produce biofuels (EU's Biofuel Directive - EC/2003/30). In 2003 the CAP Reform was entirely in force, included its limited implications in terms of price policy and market intervention; this regime remained constant over the whole 2006-2010. The only relevant change over this period has been the temporary suspension of EU import duties on cereals from January 2008 to October 2008. During the price bubble of 2007-2008, the European Union tried to curbe the cereal market turmoil by suspending the import duties for cereals though, already set at very low levels. The suspension that began in January 2008, was originally intended to last to the end of the campaign, but was postponed to June 2009; finally, the reintroduction of duties was anticipated at the end of October 2008. This change was justified by the price movements themselves (Esposti and Listorti, 2010). This temporary measure might have altered the price transmission mechanisms between international (Rotterdam) and national prices (Listorti, 2009).

Historically, the time series of agricultural markets have shown a close movement with a moderately unvaried pattern, characterized by cyclical movements with non-linear trend components in short time; whilst over the long run considerable deviations could be observed.
In recent times the increased volatility, observed from figure 1, seems to be determined by growing interactions among agricultural markets with exogenous interference of energy markets (Sumner, 2009).

Figure 1. Index of current prices of some agri commodities and oil (period 1999-2012)



Source: our elaborations. cit, wit, sit, cus, wus, sus: €/ton; for oil: €/barrel

Oil price behaviors could have affected the efficiency of agricultural market and substitution of market fundamentals (demand-supply) with other reference signals (Headey and Fan, 2008; FAO, 2009). The integration between agricultural commodity and energy market is a hypothesis to be explored, suggesting to policy makers adequate policy to mitigate the price fluctuations (Tyner and Taheripour, 2008).
These preliminary observations, suggest an hypothesis to be tested: is the oil price an exogenous price signal able to lead other markets and responsible of the increased volatility in the last years? If this is supported by some evidences it is needed a government intervention policy to re-establish the condition of market efficiency.
The time series analysis has been frequently used to study the efficiency of agricultural market (Tomek and Myers, 1993; Rosa, 1999; Thompson *et al.*, 2002; Gutierrez *et al.*, 2007; Listorti, 2009, Nazlioglu, 2011). Our analysis is developed as it follows: in the first part stationarity conditions with and without structural break are verified, in the second part the cointegration among price series is explored, in the third part the causality (linear and non linear) is analysed, and in the last part are reported the discussion of the results, main conclusions and suggestion to cope with the increased uncertain decision environment in farm planning.

2. Data collection and analysis

Data base: We considered the Italian and the US ag-commodity markets to capture affinity and/or discrepancy; specifically we focused on the soft wheat, corn and soybeans because they are the most important agricultural commodities used for feedstock, food and fuel.

The Italian time series prices of national hybrid corn (**cit**), good mercantile wheat (**wit**) and soybeans with 14% of moisture (**sit**) are obtained from DATIMA provided by ISMEA[2]; they are weekly average of daily quotations and refer to the market at the origin. The American prices are provided by FAO sourced from USDA. Corn is No. 2, Yellow U.S. Gulf (**cus**), wheat is No. 2 Soft Red Winter US Gulf Prices (**wus**) and soybeans are No.1, Yellow U.S. Gulf (**sus**). All data are weekly price series, monitored for a period spanning from January 1999 to May 2012 for a total of 699 observations, It has been decided to start with 1999 because prices since January 1999 have been reported in euro.

The fuel prices are given by weekly spot prices of brent crude oil (the North Sea BFOE complex, commonly and historically named after its first constituent crude oil, (and henceforth **oil**) provided by EIA[3]. Prices in $ are converted in € by taking the official $/€ exchange rate[4] and considering the weekly averages of daily quotation.

Table 1 presents the Pearson correlation coefficients among the price series signaling, the positive linear correlation between the oil and agricultural prices suggesting the co-movements between the price series. The high correlation coefficients of the agricultural commodity prices are consistent with the high degree of integration among agricultural markets, showing that a shock to an agri-commodity price may have an influence on other agri-commodities prices

Tab. 1. Pearson correlation coefficients[5], 1999-2012

Commodity	oil	cus	sus	wus	cit	sit	wit
oil	1.000	0.796	0.769	0.735	0.680	0.802	0.655
cus		1.000	0.885	0.794	0.770	0.798	0.782
sus			1.000	0.793	0.755	0.928	0.750
wus				1.000	0.814	0.814	0.876
cit					1.000	0.746	0.919
sit						1.000	0.755
wit							1.000
# obs							699

A comparison across periods (*pre*-price spike and *post*-price spike) reported in table 2 indicates that energy and agricultural markets became more interconnected in the more recent period of

[2] Datima is a collection of statistical databases including foreign trade and agricultural market data
[3] Energy Information Administration, Independent Statistics and Analysis
[4] Available at http://www.statistics.dnb.nl/index.cgi?lang=uk&todo=Koersen
[5] Values between 0 and 0.3 (0 and -0.3) indicate a weak positive (negative) linear relationship via a shaky linear rule. Values between 0.3 and 0.7 (0.3 and -0.7) indicate a moderate positive (negative) linear relationship via a fuzzy-firm linear rule. Values between 0.7 and 1.0 (-0.7 and -1.0) indicate a strong positive (negative) linear relationship via a firm linear rule.

observation with positive correlation for all markets in the period ‘07-’12. However, higher (or lower) values of linear correlation coefficients do not necessarily imply existence (or not existence) of causal nexus between two variables

Tab. 2. Pearson correlation coefficients, sub-periods 1999-2006 and 2007-2012

	1999-2006							2007-2012						
	oil	cus	sus	wus	cit	sit	wit	oil	cus	sus	wus	cit	sit	wit
oil	1.000	-0.150	-0.100	0.180	-0.227	0.176	-0.367	1.000	0.806	0.757	0.564	0.719	0.674	0.644
cus		1.000	0.506	0.678	0.3146	0.373	0.527		1.000	0.813	0.554	0.658	0.538	0.614
sus			1.000	0.428	0.534	0.818	0.499			1.000	0.584	0.582	0.801	0.540
wus				1.000	0.124	0.495	0.247				1.000	0.826	0.601	0.905
cit					1.000	0.400	0.616					1.000	0.610	0.927
sit						1.000	0.364						1.000	0.590
wit							1.000							1.000
# obs							417							282

The table 3 reports descriptive statistics of the price returns in each market (multiplied by 100). The returns are defined as $r_t = log(P_t/P_{t-1})$, where P_t is the price of the commodities at week t and P_{t-1} is the previous week’s price, the logarithmic transformation is a good approximation for net returns in a given market and is usually applied in empirical finance to obtain a convenience support for the distribution of the error terms. The advantage of looking at log returns of a series is that it can be observed the relative changes in a given variables and compare with other variables whose values may have very different base values.

The magnitude of the oil returns is approximately 2.5-3 times bigger than the returns in US markets and up to 6.5 times bigger than the returns in Italian corn. The oil weekly return mean is 0.14 versus 0.05 in US market (as an average of three commodities) and 0.034 in Italian market (average of Italian corn, soybeans and wheat). Oil and US agri-commodity returns have a standard deviation double bigger than the Italian returns meaning that these market are more volatile.

Tab. 3. Summary statistics for weekly returns

	oil	cus	sus	wus	cit	sit	wit
Mean	0.138	0.057	0.051	0.052	0.021	0.054	0.027
Median	0.198	0.036	0.147	-0.014	0.000	0.000	0.021
Maximum	10.192	19.547	16.883	10.055	5.577	4.557	3.996
Minimum	-9.731	-20.664	-19.337	-10.813	-7.653	-6.376	-7.918
Std. Dev.	1.951	2.092	2.093	1.988	1.073	1.002	1.022
Skewness	-0.505	-0.175	-0.688	-0.144	-0.931	-0.531	-1.997
Kurtosis	6.066	29.961	24.602	5.533	11.322	7.986	18.067
Jarque-Bera	303.148	21143.660	13626.910	188.960	2115.008	755.789	7066.649
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000
# observations	698	698	698	698	698	698	698

The returns in the seven markets appear to follow a non-normal distribution. The Jarque-Bera statistics suggest to reject the null hypothesis that the returns are well approximated by a normal distribution[6]. The kurtosis in all markets also exceeds three, pointing to a leptokurtic distribution.

3. Empirical analysis and results

3.1 - Testing for stationarity

The stationarity must be achieved before further proceeding with the analysis; this condition is firstly checked with the conventional Augmented Dicky-Fuller - ADF (1979) test that relates to the null hypothesis of non-stationarity against the alternative hypothesis of stationarity. Phillips and Perron (1988) proposed an alternative (nonparametric) method for controlling serial correlation when testing for a unit root. For comparison purposes, we performed the PP test (and reported its result) in addition to the generally favored ADF test. Rejection of the null hypothesis of a unit root under both tests is taken as evidence of stationarity. We also performed the KPSS[7] unit root test, that assumes the null hypothesis for stationary. A common problem with the conventional unit root tests, is that they do not allow for any break in the data generation process. In the case of a structural break, unit root test without structural break may result in misleading inferences. Assuming the time of the break as an exogenous phenomenon, Perron (1989) showed that the power to reject a unit root decreases when the stationary alternative is true and a structural break is ignored. This is the reason why we compared the results from conventional tests with those obtained with the Zivot and Andrews (1992) test, an endogenous structural break test which utilizes the full sample and uses a different dummy variable for each possible break date. The break date is selected where the t-statistic from the ADF test of unit root is at a minimum (most negative). Consequently a break date will be chosen where the evidence is least favorable for the unit root null. ZA test is a variation of Perron's original test with the endogenous implementation of structural breaks in the analysis: the date of the break is determined on the basis of t-statistics test of the unit root, with respect to the criteria of minimum values. Following Perron's characterization of the form of structural break, Zivot and Andrews proceed with three models to test for a unit root: model A, which permits a one-time change in the level of the series; model B, which allows for a one-time change in the slope of the trend function, and model C, which combines one-time changes in the

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[6]The **Jarque–Bera test** is a goodness-of-fit measure of departure from normality, based on the sample kurtosis and skewness. The test statistic *JB* is defined as

$$JB = \frac{n}{6}\left(S^2 + \frac{1}{4}K^2\right)$$

where n is the number of observations (or degrees of freedom in general); S is the measure of skewness (third moment), and K is the kurtosis (fourth moment) of the data. The statistic *JB* has an asymptotic chi-square distribution with two degrees of freedom and can be used to test the null hypothesis that the data are from a normal distribution.

[7] Kwiatkowski, Phillips, Schmidt and Shin (1992)

level and the slope of the trend function of the series. Hence, to test for a unit root against the alternative of a one-time structural break, Zivot and Andrews use the following regression equations corresponding to the above three models.

$$\Delta y_t = c + \alpha y_{t-1} + \beta t + \gamma DU_t + \sum_{j=1}^{k} d_j \Delta y_{t-j} + \varepsilon_t \qquad \text{(Model A)}$$

$$\Delta y_t = c + \alpha y_{t-1} + \beta t + \theta DT_t + \sum_{j=1}^{k} d_j \Delta y_{t-j} + \varepsilon_t \qquad \text{(Model B)}$$

$$\Delta y_t = c + \alpha y_{t-1} + \beta t + \theta DU_t + \gamma DT_t + \sum_{j=1}^{k} d_j \Delta y_{t-j} + \varepsilon_t \qquad \text{(Model C)}$$

where DU_t is an indicator dummy variable for a mean shift occurring at each possible break-date while DT_t is corresponding trend shift variable. The null hypothesis in all the three models is $\alpha=0$, which implies that the series $\{y_t\}$ contains a unit root with a drift that excludes any structural break, while the alternative hypothesis $\alpha<0$ implies that the series is a trend-stationary process with a one-time break occurring at an unknown point in time. The Zivot and Andrews method regards every point as a potential break-date and runs a regression for every possible break-date sequentially. Perron suggested that most economic time series can be adequately modeled using either model A or model C.

Table 4. Unit root test for series on level

		No Break			One Break	
		ADF	PP	KPSS	Zivot & Andrews	Break date
					Model A	
Constant	cus	-0.31	-1.28	1.93***	-3.63	n.a.
	sus	-1.42	-1.35	2.20***	-4.98**	July 16, 2004
	wus	-2.23	-2.15	1.79***	-3.78	n.a.
	cit	-1.77	-2.08	1.18***	-3.20	n.a.
	sit	-1.28	-1.02	2.22***	-4.03	n.a.
	wit	-1.48	-1.77	1.03***	-3.18	n.a.
	oil	-1.23	-1.17	2.48***	-3.27	n.a.
					Model C	
Constant +trend	cus	-1.59	-2.91	0.44***	-3.42	n.a.
	sus	-3.06	-3.08	0.36***	-5.48***	July 16, 2004
	wus	-3.37*	-3.29*	0.12*	-3.87	n.a.
	cit	-2.47	-2.76	0.12*	-3.82	n.a.
	sit	-2.73	-2.46	0.67**	-4.03	n.a.
	wit	-2.07	-2.36	0.57***	-3.26	n.a.
	oil	-2.78	-2.71	0.14**	-5.16**	Oct 3, 2008

Schwarz Information Criterion was used to determine the optimal lags for ADF test; the bandwidth for PP test was selected with Newey-West using Bartlett kernel. The asymptotic critical value for Zivot and Andrews test are –5.34, -4.80, -4.58 at 1%, 5% and 10% levels of significance respectively for model A (breaks in level) and -5.57, -5.08 and -4.82 at 1%, 5% and 10% levels of significance for model C (breaks in level and trend). ***/**/* denote statistical significance at 1%, 5% and 10% respectively.

Results are reported in table 4. Minimun ZA t-statistics for the levels of the variables show similar results with those obtained from the unit root tests without accounting for structural breaks with the exception for oil and sus whose results suggest that we can reject the null of unit root. Oil and soybeans_us series seem to be stationary the first one with a break in October 2008 and the second one in July 2004. Even though both price series are stationary in model C, this condition is not so evident with the same strength in model A. For this reason we consider that all the variables are integrated of order one.

As underlined by Piehl *et al*., (1999), knowledge of break point is a central point in the accurate evaluation of any program intended to bring about structural changes; such as the tax reforms, banking sector reforms and regime shifts etc.

3.2 - Cointegration analysis

The increasing co-movements between oil and agri-commodity prices during the recent years suggest to consider the cointegration relationship among the variables under investigation. With cointegrated variables it is possible to develop the relationship between the differences consistent with stationary hypothesis of the error term, on which the regression can be applied. If two or more series are individually integrated but some linear combination of them have occurred at a lower order of integration, then the series are cointegrated. The cointegration relationship can be intuitively thought like the existence of stochastic trends (such as *random walk*) common to time series. Two cointegrated time series can only diverge from an equilibrium relationship for a short period. If two time series variables are non stationary but cointegrated, at any point in time, the two variables may drift apart but there will be always a tendency for them to retain a reasonable proximity to each other.

To test for the cointegration between brent and the agricultural commodities, we performed the multivariate cointegration approach based on maximum likelihood principle suggested by Johansen (1988). The Johansen test is a procedure for testing cointegration of several I(1) time series. There are two types Johansen test, either with trace or with eigenvalue.

The trace statistic reports the null hypothesis of r cointegrated relations against the alternative of k cointegrating relations, where k is the number of endogenous variables.

The maximum eigenvalue test, on the other hand, tests the null hypothesis of r cointegrating vectors against the alternative hypothesis of $r + 1$ cointegrating vectors. The rank is calculated with the eigenvalues of a matrix. If all the eigenvalues are significantly different from zero, all processes are stationary. On the contrary, if there is at least one eigenvalue equal to zero, the process yt is integrated. On the other side, if none eigenvalue is significantly different from zero, not only the process yt is non stationary but this is for all the linear combinations. In other words there is no evidence of cointegration.

This test is not strong enough to capture the impact of structural breaks and the cointegration test over the full period may influence us towards the misleading interpretation of the result. This is the reason why, besides testing the cointegration on the full sample, we consider also some sub-periods extrapolated from the results of Z-A test: 1999w1-2004w29, 2004w30-2008w40 and 2008w41-2012w21. The Schwarz Information Criteria was used to determine the optimal lag length in VAR model to conduct the cointegration test

Table 5. Cointegration test without structural break level data

		1999w1-2012w21	**1999w1-2004w29**	**2004w30-2008w40**	**2008w41-2012w21**
		Trace statistics	Trace statistics	Trace statistics	Trace statistics
cus-oil	H_0:r=0	15.48*	19.80**	9.46	10.72
	H_0:r=1	0.79	9.55***	0.87	0.45
sus-oil	H_0:r=0	11.08	18.61**	10.30	5.22
	H_0:r=1	0.67	5.43**	1.40	0.18
wus-oil	H_0:r=0	15.42*	17.47**	8.38	6.32
	H_0:r=1	1.12	6.17**	2.24	0.72
cit-oil	H_0:r=0	13.91*	18.10**	8.55	4.21
	H_0:r=1	1.30	6.24**	1.55	0.76
sit-oil	H_0:r=0	9.55	18.00**	10.45	8.52
	H_0:r=1	0.95	3.54*	0.64	0.49
wit-oil	H_0:r=0	11.02	16.68**	13.12	4.41
	H_0:r=1	0.78	4.90**	1.58	0.35
cit-cus	H_0:r=0	20.16***	15.32*	14.83*	13.90*
	H_0:r=1	1.58	5.55**	2.33	1.68
sit-sus	H_0:r=0	20.01***	27.91***	18.76***	15.31*
	H_0:r=1	1.58	3.22*	0.71	1.67
wit-wus	H_0:r=0	27.97***	11.85	27.49***	27.33***
	H_0:r=1	3.54	4.14*	1.53	1.73

p-values are MacKinnon-Haug-Michelis (1999) *p*-values; ***/**/* denote rejection of the null hypothesis at the 1%, 5% and 10% level.

The first part of table 5 reports the results of the cointegration test between brent and agri-commodities in the US market for the entire 1999-2012 period and sub-periods. The results indicate that there are similar cointegration relations for cus-oil and wus-oil prices The best results of cointegration are obtained in the sub-period 1999-2004 for all the commodities; by observing the price trends this is the less turbulent period during which we hypothesize the absence of antagonist fluctuation that disturb the market integration. Some studies, on the other hand, report varying results based on different time periods. For example, Campiche *et al.* (2007) examined the co-movements between world crude oil prices and corn, sorghum, sugar, soybeans, soybean oil, and palm oil prices during the period 2003–2007 based on weekly data. The empirical analysis with the Johansen cointegration test shows that while there is no cointegrating relation among the variables in concern for the period 2003–2005, corn and soybean prices are cointegrated with crude oil prices during the period 2006–2007. Harri *et al.* (2009) report a consistent cointegrating relationship between crude oil and corn, soybeans starting in April 2006. Nazlioglu (2011) considers the cointegration between oil and the three key agri-commodity prices (corn, soybeans and wheat) and reports that corn and soybeans are cointegrated with the oil prices during the period 2008-2010. our results differ from those found in literature mostly because of the different length of the sub-samples considered.

The second part of table 5 reports the cointegration results for brent and agri-commodity prices in Italy. In this case we observe some similar behaviour with the American market and oil in the first sub-period, and absence of cointegration in all the other sub-periods.
Such results suggest to further investigate the cointegration among the American and the Italian commodity markets that is a hypothesis supported by the Law of One Price. This economic principle states that the same item or closely equivalent items must be sold for the same price or related prices in the marketplace; in practice prices move together, they may not exactly match but could be pretty close, and market forces affect the whole market. Following the usual definition, commodity arbitrage ensures that each good has a single price (defined in a common currency unit) throughout the world (Isard, 1977).
The results of this further analysis are reported in the third part of table 5 and suggest clear evidence of cointegration that is statistically significative for the entire period of observation and in the sub-periods. This support the idea that volatility in Italian agri-commodity market is induced by US market.
Following the methodology proposed by Nazlioglu (2011), to be sure to give a good interpretation of the results from Johansen's (1988) testing framework, since the structural break dates were determined *a priori* instead of finding them endogenously in the cointegration model, the relationship between brent and ag-commodity prices was also analyzed by running the Gregory-Hansen program. This is a residual-based test for cointegration in models with regime shift proposed by Gregory and Hansen in 1996. They examine tests for cointegration which allow the possibility of regime shift and develop ADF^*, Z_t^* and Z_α^* type tests designed to test the null of no cointegration against the alternative of cointegration in the presence of a possible structural break that can occur in intercept (level shift, model C), in intercept with trend (level shift with trend, model C/T) or in cointegration vector (regime shift, model C/S). The structural change is endogenously determined by the smallest value (the largest negative value) of the cointegration test statistics across all possible break point.
Table 6 shows the results of cointegration between oil and US agri-commodity prices. As far as brent and cus price relation is concerned, ADF* fails to reject the null hypothesis of no cointegration with model C, C/T and C/S whereas Z_t^* and Z_α^* type test results indicate the rejection of the null for all the three models with break dates in 2004 and 2010. For the case of soybeans and brent, all three tests do not reject the null hypothesis of cointegration presenting a structural break in August 2004; besides, Z_t^* and Z_α^* fail to reject the null in the regime shift model with a break in July 2007. For the long run relationship between wheat and brent prices, the are no evidence of a cointegration relationship. These results confirm those found by Harri *et al.* (2009) where the presence of cointegrating relations between crude oil and agricultural commodities (corn, soybeans,

soybean oil, cotton and wheat) is verified. The authors conclude that in the most recent periods no cointegrating relation exists between crude oil and wheat, likely because wheat prices have been heavily influenced by weather events, meaning less influence of input prices (as well as the fact that wheat is not significantly used in ethanol production).

Table 6. Cointegration test with structural break (Gregory Hansen test) between US agri commodities and oil

		cus-oil		sus-oil		wus-oil	
US-oil		test stat	break date	test stat	break date	test stat	break date
ADF*	C	-3.45	n.a.	-4.21	n.a.	-4.23	n.a.
	C/T	-3.84	n.a.	-5.38**	Aug 20, 2004	-4.26	n.a.
	C/S	-4.06	n.a.	-4.94*	Mar 7, 2008	-4.65	n.a.
Z_t^*	C	-4.69**	May 14, 2010	-4.44*	Sept 28, 2007	-3.81	n.a.
	C/T	-5.52***	May 28, 2004	-5.72***	Aug 13, 2004	-3.85	n.a.
	C/S	-5.72***	Sept 10, 2004	-5.20**	Sept 28, 2007	-4.03	n.a.
Z_α^*	C	-42.20**	May 14, 2010	-40.26**	Sept 28, 2007	-28.61	n.a.
	C/T	-56.96**	May 28, 2004	-61.15***	Aug 13, 2004	-28.91	n.a.
	C/S	-62.39***	Sept 10, 2004	-52.52**	Sept 28, 2007	-31.49	n.a.

Model C: Level shift, Model C/T: level shift with trend, Model C/S: Regime shift. Null hypothesis: no cointegration. For ADF* and Z_t^* tests, critical values in Model C are: -5.13 at 1%, -4.61 at 5% and -4.34 at 10%; in Model C/T:-5.45 at 1%, -4.99 at 5% and -4.72 at 10%; in Model C/S: -5.47 at 1%, -4.95 at 5% and -4.68 at 10%. Critical values for Z_α^* test are -50.07, 40.48, -36.19 respectively at 1, 5 and 10% in Model C; -57.28, -47.96 and –43.22 at 1, 5 and 10% in Model C/T; -57.17, -47.04 and -41.85 at 1, 5 and 10% in Model C/S. The optimal lag length for ADF* test was selected by Akaike information criterion. ***/**/* denote statistical significance at 1%, 5% and 10% level of significance, respectively.

In table 7, the tests do not support the evidence of cointegration among the brent prices and the Italian commodity prices confirming the result obtained with the Johansen test.

Table 7. Cointegration test with structural break (Gregory Hansen test) between IT agri commodities and oil

		cit-oil		sit-oil		wit-oil	
IT-oil		test stat	break date	test stat	break date	test stat	break date
ADF*	C	-3.77	n.a.	-4.11	n.a.	-3.75	n.a.
	C/T	-4.01	n.a.	-4.13	n.a.	-3.84	n.a.
	C/S	-4.27	n.a.	-4.20	n.a.	-4.27	n.a.
Z_t^*	C	-3.58	n.a.	-3.69	n.a.	-3.36	n.a.
	C/T	-3.58	n.a.	-3.71	n.a.	-3.33	n.a.
	C/S	-3.73	n.a.	-3.84	n.a.	-3.59	n.a.
Z_α^*	C	-24.72	n.a.	-27.95	n.a.	-21.93	n.a.
	C/T	-24.63	n.a.	-28.21	n.a.	-22.15	n.a.
	C/S	-28.14	n.a.	-29.42	n.a.	-25.67	n.a.

The results of Gregory-Hansen test reported in table 8 clearly emphasize the existence of cointegration relationships between the Italian and the American commodity markets. In the case of wheat and soybeans there is a strong evidence of such relation for all the tests confirming the results obtained in the cointegration test without structural breaks, whereas in the case of corn evidence of cointegration is supported by Z_t^* and Z_α^* tests.

Table 8. Cointegration test with structural break (Gregory Hansen test) between It and US agri commodities

		cit-cus		sit-sus		wit-wus	
IT-US		test stat	break date	test stat	break date	test stat	break date
ADF*	C	-3.89	n.a.	-4.83**	May 14, 2010	-5.29**	Ago 24. 2001
	C/T	-4.21	n.a.	-4.96*	May 14, 2010	-5.26**	Apr 6 2001
	C/S	-4.43	n.a.	-6.11***	July 18, 2008	-5.56***	July 9, 2004
$\mathbf{Z_t^*}$	C	-4.81**	Aug 1, 2008	-7.11***	May 14, 2010	-5.70***	May 11, 2001
	C/T	-5.10**	July 4, 2003	-7.04***	May 14, 2010	-5.71***	May 11, 2001
	C/S	-4.86**	June 27, 2008	-7.63***	May 14, 2010	-5.98***	July 16, 2004
$\mathbf{Z_\alpha^*}$	C	-45.21**	Aug 1, 2008	-90.15***	May 14, 2010	-61.46***	May 11, 2001
	C/T	-50.52**	July 4, 2003	-88.94***	May 14, 2010	-61.49***	May 11, 2001
	C/S	-46.00*	June 27, 2008	-103.52***	May 14, 2010	-67.48***	July 16, 2004

Since the cointegration analysis cannot determine the direction of causality, it is common to investigate causal interactions among the variables once cointegration is established.

Results obtained from the previous statistical analysis, suggest us to go further into the investigation of linear-causal dynamics by looking for the direction of the causality between the observed series. To formally analyze the dynamic relationship between brent and agri-commodity prices, linear Granger causality test was conducted. This test allows us to examine whether changes in the price of oil leads changes in agricultural prices. The idea was to make some inferences about the direction of information flows among markets.

A time series *X* is said to Granger-cause *Y* if it can be demonstrated with a series of t-tests and F-tests on lagged values of *X* (and with lagged values of *Y* also included), that those *X* values provide statistically significant information about future values of *Y*. The null hypothesis is that *X* does not Granger cause *Y*. If the null is rejected, then, there is causality.

This test, as the Johansen cointegration test, is not able to capture the impact of structural breaks and it may lead us towards the misleading interpretation of the result. For this reason linear causality tests were performed over the entire sample period, as well as on sample subperiods, to analyze whether the dynamic relationship between oil and ag-commodities prices has changed across time.

Table 9 reports Granger causality-test results along with some diagnostic test. The Durbin Watson test tat is used to detect the presence of autocorrelation in the residuals from a regression analysis, implies that the residual of most of the estimated model are free from autocorrelation problems (apart from the models obtained from the regression between oil and Italian agri commodities whose DW statistic results are substantially less than 2, with a clear evidence of positive serial correlation). The Ramsey Regression Equation Specification Error Test (RESET) test that is a general specification test for the linear regression model, clearly shows that the functional forms for the models are appropriately specifies (with some exception in the comparison between oil and Italian commodities)

Table 9. Linear Granger causality test

	1999w1-2012w21		1999w1-2004w29		2004w30-2008w40		2008w41-2012w21	
	F-Statistic	**Probability**	**F-Statistic**	**Probability**	**F-Statistic**	**Probability**	**F-Statistic**	**Probability**
oil →cus	0.71	0.55	1.76	0.18	2.65	0.07	0.25	0.61
cus →oil	0.92	0.43	1.05	0.35	0.35	0.71	1.01	0.32
Durbin-Watson	2.22		2.15		2.11		2.28	
RESET	0.08	0.78	1.47	0.23	1.13	0.29	0.12	0.73
oil →sus	0.18	0.67	1.24	0.27	0.69	0.50	3.42	0.03
sus →oil	0.38	0.53	0.83	0.36	0.17	0.84	1.46	0.23
Durbin-Watson	2.27		2.54		1.88		2.25	
RESET	0.24	0.63	0.00	0.97	0.99	0.32	0.00	0.99
oil →wus	1.45	0.23	0.77	0.38	0.08	0.78	5.03	0.01
wus →oil	0.77	0.38	1.29	0.26	0.61	0.44	0.51	0.60
Durbin-Watson	2.14		2.04		1.97		2.12	
RESET	0.90	0.34	0.80	0.37	0.09	0.76	0.50	0.48

	1999w1-2012w21		1999w1-2004w29		2004w30-2008w40		2008w41-2012w21	
	F-Statistic	**Probability**	**F-Statistic**	**Probability**	**F-Statistic**	**Probability**	**F-Statistic**	**Probability**
oil →cit	2.08	0.15	0.17	0.68	0.00	0.97	0.24	0.79
cit →oil	0.34	0.56	1.71	0.19	1.19	0.28	0.07	0.93
Durbin-Watson	1.14		1.43		1.32		1.07	
RESET	4.50	0.03	0.66	0.42	0.22	0.64	2.00	0.16
oil →sit	2.13	0.14	0.14	0.71	0.01	0.93	0.64	0.53
sit →oil	1.15	0.28	0.66	0.42	0.08	0.77	1.02	0.36
Durbin-Watson	1.61		1.43		1.67		1.75	
RESET	11.15	0.00	0.30	0.59	0.48	0.49	3.11	0.08
oil →wit	0.45	0.64	0.32	0.57	2.36	0.13	0.83	0.43
wit →oil	0.02	0.98	0.03	0.86	1.61	0.21	1.19	0.31
Durbin-Watson	1.33		1.54		1.19		1.57	
RESET	2.57	0.11	1.25	0.27	0.21	0.66	3.05	0.08

	1999w1-2012w21		1999w1-2004w29		2004w30-2008w40		2008w41-2012w21	
	F-Statistic	**Probability**	**F-Statistic**	**Probability**	**F-Statistic**	**Probability**	**F-Statistic**	**Probability**
cus → cit	2.36**	0.04	0.09*	0.77	3.72**	0.03	1.77*	0.07
cit → cus	0.56	0.72	1.23	0.27	0.64	0.53	0.55	0.85
Durbin-Watson	1.14		1.40		2.08		2.34	
RESET	0.01	0.92	0.89	0.45	2.51	0.12	1.01	0.32
sus →sit	4.49***	0.00	4.82**	0.03	14.89***	0.00	16.54***	0.00
sit →sus	0.19	0.31	3.34*	0.07	0.83	0.44	1.99	0.14
Durbin-Watson	1.76		1.46		1.90		2.01	
RESET	0.30	0.58	0.52	0.47	0.14	0.71	0.71	0.67
wus → wit	14.10***	0.00	0.44	0.51	4.61***	0.00	7.44***	0.01
wit →wus	8.64***	0.00	0.01	0.93	1.86	0.12	1.68	0.20
Durbin-Watson	1.42		1.52		1.35		1.62	
RESET	1.92	0.17	0.05	0.82	2.41	0.12	0.01	0.92

→means non Granger causality hypothesis. ***/**/* denote statistical significance at 1%, 5% and 10% level, respectively. The optimal lag length was selected by Schwarz information criterion.

The upper section of the table reports the F-statistic for the null hypothesis of no causality between brent prices and the US agricultural market (and *vice versa*); results show that there is no causal linkage from the oil prices to the agricultural prices. The linear Granger causality analysis supports the neutrality hypothesis in the full period of observation an in the sub-samples these results are consistent with the findings of Nazlioglu (2011).

The middle section reports the results for the Granger Causality analysis between brent and the Italian commodities. In general it can be said that there is no linear causality between oil prices to the agri-commodity prices which is consistent with what Zhang and Reed (2008) found for the Chinese grains.

More clear appear the linear linkages among the agricultural commodities in the two markets reported in the third section of the table. In all three cases the US ag-commodities do Granger-cause the Italian prices in the entire sample period and sub-samples as expected by the low of one price.

Some authors argue that the traditional Granger causality test, designed to detect linear causality, is ineffective in uncovering certain nonlinear causal relations and recommend the use of nonlinear causality tests, Baek and Brock (1992), Hiemstra and Jones (1994).

Considering, then, that linear causality tests might overlook nonlinear dynamic relations between brent and ag-commodities, according to Nazlioglu (2011), we conducted the nonparametric causality test proposed by Diks and Panchenko (2006, hereafter DP) which avoids the over-rejection observed in the test proposed by Hiemstra and Jones (1994). The DP test detects nonlinear causal relationships with high power, but does not provide any guidance regarding the source of the nonlinear dependence. Identifying a specific nonlinear dynamic relationship between oil and agricultural markets is beyond the scope of the present study. The DP analysis was carried out in two steps: it was first applied to stationary series, and then, to remove any linear dependence, the test was applied to the estimated residual series from the VAR model with the pair of variables of interest. "By removing linear predictive power with a linear VAR model, any remaining incremental predictive power of one residual series for another can be consider non linear predictive power" (Hiemstra and Jones, 1994). The tests were performed for different lag values.

Following Dicks and Panchenko (2006), the data were also normalized to unit variance before the test was applied and the bandwidth, which value plays an important role on the existence of non linear causality, was set to 1, as it is one time the standard deviation. Because nonparametric tests rely on asymptotic theory, causality tests on sample subperiods were not performed in this case. Table 10 reports the T values for Diks and Panchenko's test statistic applied to the variables and to their residuals in both directions and for different lag lengths (1–2 lags).

Table 10. Nonlinear Granger causality test (Diks –Panchenko test)

	Raw data		Residuals	
lags	**1**	**2**	**1**	**2**
oil →cus	2.019**	2.788**	1.434*	2.282**
cus →oil	0.210	0.812	0.450	0.144
oil →sus	1.686**	1.826**	1.570*	1.796**
sus →oil	0.395	0.172	0.311	0.250
oil →wus	1.290*	1.282*	0.968	1.329*
wus →oil	1.969**	1.829**	2.140**	2.135**

→means non Granger causality hypothesys. ***/**/* denote statistical significance at 1%, 5% and 10% level of significance, respectively.

In the table 10 we reported the results of the non linear Granger causality test between brent and US agri-commodity prices. As regards **oil and corn**, there is evidence of unidirectional information flow for raw data that is confirmed even after filtering the series with VAR model. In terms of volatility, brent prices seem to transfer volatility to corn prices. Our findings are consistent with those obtained by Nazlioglu (2011). Until early 2007, corn prices had little relationship to crude oil prices. Since then, however, corn prices have been more responsive to changes in crude oil prices. The main reason for this is the growing importance of fuel ethanol as a percentage of total demand for US corn. With the current large size of the ethanol industry, corn prices have become closely related to crude petroleum prices because corn is now a major energy crop. With rapid growth of the ethanol industry in the last few years, corn has become very much an energy crop as well as the world's most important source of feed grains for production of livestock, poultry, and dairy products. This transition has created a strong but still somewhat variable relationship between corn prices and those for crude oil and ethanol. Ethanol, has a price relationship with crude oil. This relationship has fluctuated in a moderate range in the past few years. At times when ethanol demand and prices have surged in the past, corn prices have tended to be strengthened.

In the case of **brent and soybeans**, the non linear causal test on raw data shows a direct relationship between the pair of variables of interest. Besides, after removing any linear dependence, the results of the VAR residuals, at lag one and two, support the existence of a strict nonlinear price transmission from the brent to the soybeans prices which persists over the long run. Even in this case, then, the crude oil market has a strong influence on the agri-commodity prices, since it influences biofuel market, and soybeans are still the dominant feedstock for biodiesel.

For the causality between **brent and wheat**, the nonparametric results provide a feed-back evidence in the relationship between the variables prices (raw data). Further investigation with the VAR residuals indicates that the causal relationship goes in the reverse direction at lag one and in a bi-directional way at lag two. These findings remark the different behaviour of the wheat with respect to corn and soybeans related to its application mostly to the industrial sector. Besides, apart from brent influence, there are other exogenous factors influencing the price volatility of wheat. The weather is the most important factor of influence on the supply situation for agricultural products. Dry weather and long persistent drought, foe example, can have great impact on supplies and prices, as well as extremely wet periods, extreme temperature fluctuations or environmental disasters. Further important influence factors are: the seasonal growing and harvesting rhythm (price-lows often in the harvesting months in summer), and the influence from the side of politics.

In this paper we do not report the results for the nonlinear causality analysis of oil/Italian commodities and IT/US agri commodities (results upon request). In general there isn't a non linear causality between brent and Italian agri-commodity prices and these results are consistent with those obtained by Rosa and Vasciaveo (2012). In the case of non linear causality between the US and the Italian commodity prices, raw data provide a feed-back evidence in the relationship between the variables prices and the same findings are presented after filtering the series. Such misleading results could be explained with the fact that there is a one way strictly *linear causality* among the examined variables (Table 9)

4. Conclusion

The main goal of our work was to observe whether the increasing volatility in crude oil prices could cause higher price changes in selected agricultural commodity markets. In order to find the causal linkage among the variables under investigation, we used both linear and non linear Granger causality methods.

The results of the linear Granger causality analysis suggest to accept the presence of neutrality hypothesis in the US markets which means that the prices of oil and the agricultural commodities do not cause each other in a strictly linear sense. Same results are evident in the Italian market. In Italy the prices are co-integrated with US corresponding prices of agri-commodities and there is also evidence of linear unidirectional Granger causality from US to Italian market. In Italy co-integration and causality among commodities suggest that the volatility is transmitted from global to local agri-commodity prices and the local agricultural commodity prices do not respond to the world oil prices. These results are confirmed by the Law of One Prices.

Diks Panchenko test provide clear evidence of the strong non linear relationship between oil and corn, which is indicative of the growing use of corn for ethanol. The crude oil is also a *leading indicator* for soybeans, it is exogenous to the formation of its prices and able to affect their realization. The brent volatility contributes to destabilize the prices of corn and soybeans because of their energy and industrial use. The case of wheat prices is completely different: the world has consumed more wheat than has been produced in six of the last seven years. The resulting drawdown in wheat stocks is largely responsible for the large increase in wheat prices. This perception of food insecurity, due to the diminishing supply of flours, has brought wheat prices to surge upwards dramatically for the financial speculation prevailing on market fundamentals.

Volatility becomes an issue for policy response when it induces risk averse behaviour that leads to inefficient investment decisions and when it creates problems that are beyond the capacity of producers, consumers or nations to cope with. To be effective the market policies need to have unbiased information about the agri-food supply chain, producers, consumers and traders to reduce the risk of market inefficiency. It is necessary to focus on the policy options designed to prevent or reduce price volatility and mitigate its consequences: some would help to avert a threat, others are in the nature of contingency plans to improve readiness, while still others address long-term issues of resilience.

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